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Informational efficiency of the US SO_2 permit market

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Abstract

We test the information efficiency of the market for SO_2 permits in the US. In order to do so, we analyse the price process of these permits using techniques that have been widely used in financial economics. The evidence presented in this paper suggests that this market is efficient from an informational point of view. Although one could question this hypothesis from a statistical point of view, economic significance suggests that this market is indeed efficient.

Keywords: SO_2; Tradable permits market; Market efficiency; Unit roots; Random walk hypothesis

1. Introduction

The sulphur dioxide (SO_2) regulation in the United States gradually evolved from a body of technical regulation with national air quality standards and New Source Performance Standards (NSPS) for new power plants into an innovative trading program in SO_2 emissions allowances. The trading program followed from Title IV of the 1990 Clean Air Amendments that set a goal in 2010 of reducing annual SO_2 emissions by 10 million tons from the 1980 level. Phase I of the trading scheme began in 1995 and affected 263 units at 110 mostly coal-burning electric utility plants located throughout 21 eastern and Midwestern States. Phase II, which began in 2000 further tightened annual emission restrictions on the larger, higher emitting Phase I plants and set emission restrictions on smaller, cleaner plants. Participation into the program has been strong and it is generally acknowledged that the flexibility of the program provided annual cost savings of approximately \$0.9 billion to \$1.8 billion compared to costs under a command-and-control regulatory alternative (Council of Economic Advisors, 2004). With the offered flexibility, emitters have the freedom to decide how, when and which measures will be taken to lower or not SO_2 emissions.

From the point of view of environmental policy, efficiency of permit trading is a key issue that is assumed in most of the work on the way in which permits can be used to address environmental problems (Joskow et al., 1998). Indeed, without efficient markets, permit prices cannot give accurate signals to market participants. Hence, in inefficient markets, it would be hard to assume that the decision to abate and sell or not to abate and to buy permits would be efficient.

Market efficiency can be analysed from a number of different perspectives. First of all, one can look at the process of matching supply and demand. The question to be answered from this perspective is whether the SO_2 permit market resembles a competitive and frictionless market? Joskow et al. (1998) have used this approach to study the efficiency of the US SO_2 permit market. They argue that their analysis of the "*evolution*

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of the sulphur dioxide allowance market indicates that a relative efficient private market developed in a few years time, by at least mid-1994" (Joskow et al., 1998, p. 683).

A second perspective borrows from the financial economics literature and analyses the outcome of the process of matching supply and demand: the permit price. Here, the focus of the analysis is the speed with which new information works its way into permit prices. If permit markets are efficient, they should reflect new information fast. Assume, for instance, that at time t new information reaches the permit market which would drive permit prices up. In an efficient market, if the price should move up, it should do so at once and not in series of small steps (LeRoy, 1989). Hence, at time $t+1$ permit prices should have adjusted and should reflect the new information. In other words, if new information arrives, permit prices adjust immediately and hence, they continue to give accurate pricing signals to market participants. If it takes a long time before permit prices fully reflect the new information and it takes a number of periods before prices are fully adjusted, permit prices give the wrong signal to market participants as long as the adjustment process takes place. As Malkiel puts it: markets are efficient from an informational point of view "*if tomorrow's price change will reflect only tomorrow's news and will be independent of price changes today*" (Malkiel, 2003, p. 59).

In this paper we analyse the efficiency of the US SO_2 permit market from an informational point of view. If the SO_2 permit market is efficient in an informational sense, this would be evidence which supports the hypothesis that market participants have a good idea of the market-clearing price and the influence of new information on this price level. Indeed, if only today's news impacts the market price and yesterday's news does not have an influence, market participants must have a good sense of its impact on market-clearing prices. If, on the other hand, the SO_2 market is found to be inefficient and yesterday's news has an impact on today's prices, this would support the hypothesis that participants have no good sense of the market-clearing price and they are either slow to react to new information or overreact.

In terms of the analysis, informational efficiency requires that today's news does not help us to predict tomorrow's prices. If market participants are able to quickly assess the impact of new information on the permit's price, by the end of the day, today's permit price should reflect today's news. It means that a small increase in the permit's price today would inform market participants of one or more small increases tomorrow and in the days ahead as the market participants work their way through the assessment process. Examining the SO_2 permit price time series allows us to analyse if past price changes are informative with respect to future price changes. We can focus on the SO_2 permit price history and we can disregard the impact of other variables. Assume, for instance, that there is a variable x whose behaviour has an impact on permit prices. Say that a change in x causes the permit price to rise. From the point of view of our analysis, the way x behaves in a specific period is uninformative. Assume, for instance, that in period t, x rises and that this should cause a rise in the permit price equal to δ. If permit markets are efficient, this means that period t permit prices should completely reflect the increase in x and should rise with δ. If they are not efficient, the rise with δ in the permit price will take n periods as the permit prices only adjust in small steps of say δ/n. Hence, in that case the rise in the permit price in period t with δ/n would inform market participants of a number of small rises yet to come. Hence in our analysis, we do not need to know the cause of the permit price change. We can focus on the question if and to what extend are past price changes in the permit market informative for future changes to come?

As market efficiency is one of the key conditions for any permit scheme to work properly, an analysis of the US SO_2 permit market is a worthwhile exercise. The remainder of this paper is organized as follows: Section 2 discusses some theoretical issues. In Section 3, we proceed with an empirical analysis of the US SO_2 permit market. The final section concludes.

2. SO_2 permit prices as random walks

We will start our analysis from the assumption that the SO_2 permit price process can be modelled as a random walk. The idea of a random walk and the efficient market hypothesis are closely related (Malkiel, 2003). Assume for instance that the market price for a permit at time t equals P_t. If market participants are rational, the price P_t reflects the expected value of the permit given the information they have at time t. The information set at time t, I_t, contains all the data which are necessary and available to value the permit, i.e. to determine the level of P_t. I_t contains the price of the permit itself but also expectations with respect to policy changes, new technologies, future market conditions, prices for oil, electricity, etc. As new information becomes available, new data are added to the information set. Assuming that market participants have an infinitely long memory, this means that the information set at time t is contained within the information set of the next period $I_t \subset I_{t+1}$: all the data available at time t are also available in the next period $t+1$. Notice that the new information is not contained in the information set at time t. If, for instance, new technologies emerge at time $t+1$ but their arrival was widely expected at time t these new technologies do not represent new information and they are part of I_t. If, on the other hand, a policy change announced at time $t+1$ surprises market participants, it is part of I_{t+1} but not of I_t. Given I_{t+1} market participants reassess their rational expectation of the value of a permit. A new price P_{t+1} reflects this reassessment. By definition, the flow of new information is uncertain. New information arrives at random intervals in time and the impact of this information can be positive or negative for, permit prices can require a large of small reassessment. As the flow of information is random, price changes of permits are random as well.

From the previous discussion, it follows that we need a model which relates the random news flow to the changes in the permit price. Consider the following model for the natural logarithm of the price of an SO_2 permit at time t (which we will denote with $p_t = \ln(P_t)$ and we will use 'SO_2 permit price' to refer to the natural logarithm):

$$p_t = p_{t-1} + \gamma + \varepsilon_t \quad (1)$$

with γ a drift parameter and ε_t the random increment of the process with $\mathrm{E}[\varepsilon_t] = 0$ and $\mathrm{var}[\varepsilon_t] = \sigma^2$. The random increment can be seen as the impact on today's price of today's news. Eq. (1) says that the rational expectation of market participants of the value of a permit today equals their rational expectation in the previous period adjusted for the impact of unexpected news (see $\mathrm{E}[\varepsilon_t] = 0$).

Campbell et al. (1997) distinguish three types of random walks based on the properties of the increments ε_t. If they are identically and independently distributed (IID-property), Eq. (1) is a random walk of type 1 (RW1). The IID-property implies that for two arbitrary functions f and g and scalar $k \neq 0$.

$$\mathrm{cov}[f(\varepsilon_t), g(\varepsilon_{t-k})] = 0 \quad (2)$$

The increments ε_t of a type 2 random walk (RW2) are independent but not identically distributed. Type 3 random walks (RW3) are characterized by dependent but uncorrelated increments. RW3 implies that the equality in Eq. (2) holds for all linear functions f and g but not for non-linear functions.

From Eq. (1) it follows that the condition in Eq. (2) could also be written in terms of $\Delta p_t = \gamma + \varepsilon_t$.

If the SO_2 permit prices are random, the permit market is said to be (weakly) efficient as it is impossible to profit by trading on the information contained in the permit price history (Campbell et al., 1997). However, even if this is not the case, the market may still be efficient as each transaction involves trading costs. Hence, the permit market would still be efficient as long as the information contained in the price history is insufficient to allow a market participant to earn a profit after transaction costs have been accounted for. As such one has to judge whether the results are significant from a statistical as well as an economic point of view (Malkiel, 2003).

3. Efficiency of the US SO_2 permit market

3.1. Data

To analyse the behaviour of the US SO_2 permit prices we have used monthly data from August 1994 to December 2001 (89 observations). Fig. 1 shows the series based on price information by the US Environmental Protection Agency (EPA). EPA (2001) presents price data from Cantor Fitzgerald and Fieldston Publications. Fig. 1 is based on data from Fieldston Publications. We have chosen not to include the year 2002 as the SO_2 and other permit programs in the US might have been affected by the Californian energy crisis. We have done so because it is not unreasonable to assume that the temporary halting of the NO_x permit market did not have an impact on the market for SO_2 permits. Due to the havoc on the NO_x market, excessive uncertainty with respect to the future functioning of the SO_2 permit market for instance could have influenced market participants. Given the data that are available to us, including the Californian energy crisis could put to much weight on this a-typical period.

Table 1 reports the summary statistics for both p_t as well as its first difference Δp_t. Fig. 1 shows the price history of p_t.

3.2. Unit root tests

A random walk is a first difference stationary process. Hence, the first issue to be looked at is whether the SO_2 permit price series contains a unit root while the first difference of this series does not.

A number of alternative tests are available to analyse if a process contains a unit root. The Augmented Dickey Fuller (ADF) test uses the following regression:

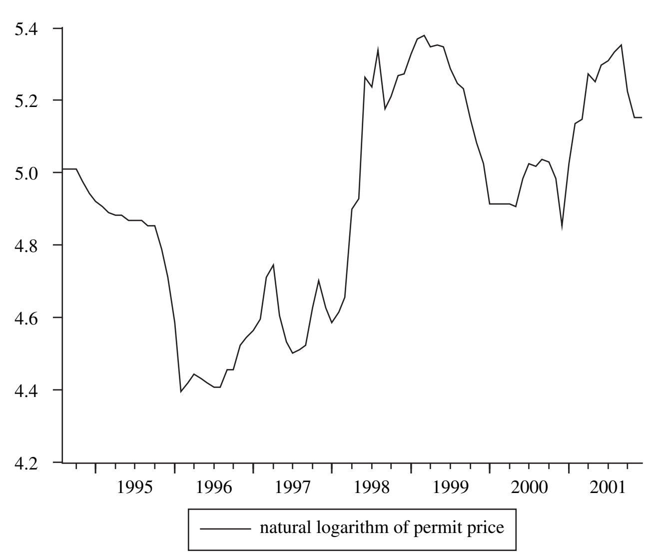

Fig. 1. Natural logarithm of the SO_2 permit price.

Table 1
Summary statistics for p_t and Δp_t

	p_t	Δp_t
Mean	4.9177	0.0034
s.e. of mean	0.0317	0.0010
t-Statistic	154.9263	3.2140
Variance	0.2994	0.0004
Skewness	−0.1806	1.7069
Kurtosis	−1.1389	11.4473
Jarque–Bera	5.2943	2063.1517
Observations	89	88

$$p_t = \alpha + \gamma t + \theta p_{t-1} + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \tag{3}$$

where α and γ are drift and deterministic trend components, respectively, – to determine if $\theta = 1$ against the alternative that $\theta < 1$. As is well known, under the null of a unit root, the test statistic (the *t*-statistic on θ) does not have a standard distribution (see e.g. Hamilton, 1994 or Verbeek, 2000).

The ADF-test requires that error terms ε_t are independent and homogeneous. The addition of lagged differences in Eq. (3) is meant to remove any autocorrelation from the error terms. The number of lags k that should be included to make the ε_t sequence white noise is, however, unknown. Various alternatives have been explored to select the optimal lag length. Model specification tests use the Akaike information criterion (AIC) or the Bayesian information criterion (BIC). However, as Ng and Perron (1995) have shown, these criteria often select very small values of k. The general to simple procedure starts with a specified number of lags k and reduces that number to $k - 1$ lags if the lag of length k is insignificant.

The test due to Phillips and Perron (1988) (PP) is a generalization of the ADF-test and is less demanding with respect to the error terms. The PP-test allows the error terms to be weakly dependent and heterogeneously distributed by including a weighting function and various lags of the error process to calculate a consistent estimate of the variance.

Both the ADF and PP's null hypothesis are that a series contains a unit root. However, unit root tests often lack power. Kwiatkowski, Phillips, Schimdt and Shin (1992) (KPSS) propose a test whose null is stationarity. Their test is based on the residuals from

$$p_t = \alpha + \gamma t + \varepsilon_t \tag{4}$$

The test statistic is given by $\sum_{t=1}^{T} S_t^2 / \hat{\sigma}^2$ with $S_t = \sum_{s=1}^{t} \varepsilon_s$ and $\hat{\sigma}^2$, an estimator of the variance of the error terms. The 5% KPSS critical value for the null of trend stationarity equals 0.146. To test the null of stationarity, the trend is omitted from Eq. (4) and the 5% critical value equals 0.463 (Kwiatkowski et al., 1992). In order to compute $\hat{\sigma}^2$, KPSS propose a procedure similar to PP and include a weighting function to correct for autocorrelation.

Perron (1997) proposes a test that includes the possibility of changes in the intercept and slope of the deterministic trend in Eq. (1). Indeed, if there is a one-time increase in the intercept of a trend-stationary process, standard unit root tests are biased towards accepting the null of a unit root. Three models, all of which use OLS estimates, are used to test if $\theta = 1$ in the presence of a break at time $t = \tau + 1$. The test statistic in all three cases is the *t*-statistic for the test that $\theta = 1$. The distribution of this *t*-statistic is non-standard. Perron (1997), however, provides critical values for various sample sizes and models.

The first model tests the null of a unit root with a one-time shift in the non-stationary process against the alternative of a trend-stationary process with a one-time shift in the intercept. The test uses the following regression:

$$p_t = \alpha + \gamma t + \psi D_L + \theta p_{t-1} + \delta D_P + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \tag{5}$$

with

$$D_L = \begin{cases} 0 \Leftrightarrow t \le \tau \\ 1 \Leftrightarrow t > \tau \end{cases} \quad \text{and} \quad D_P = \begin{cases} 0 \Leftrightarrow t \ne \tau + 1 \\ 1 \Leftrightarrow t = \tau + 1 \end{cases}$$

The second model allows for a change in both the intercept and the slope of the deterministic trend and uses

$$p_t = \alpha + \gamma t + \psi D_L + \varphi D_L t + \theta p_{t-1} + \delta D_P + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \tag{6}$$

The alternative hypothesis in Eq. (6) is a one-time change in the intercept and slope of a trend-stationary process. Finally, the third model allows for a change in the slope of a trend-stationary process but assumes that both segments of the trend are joined at the time of the break. The test uses

$$p_t = \alpha + \gamma t + \varphi D_L (t - \tau) + \theta p_{t-1} + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \tag{7}$$

to test if $\theta = 1$. Hence, the hypothesis of a unit root is tested against the alternative of a change in the slope of a trend-stationary process.

Perron (1997) proposes various alternatives to select the break data t endogenously. The first minimizes the *t*-statistics on $\theta = 1$. The second alternative minimizes the *t*-statistic on ψ (model 1) or on φ (models 2 and 3). The third alternative is similar to the second one but uses the absolute values of the *t*-statistics. To determine the lag length k, Perron (1997) proposes the general to specific procedure.

We have performed unit root tests for the SO_2 permit price series (p_t) as well as the first difference of this series (Δp_t). Table 2 reports the results.

First of all, we have used the ADF-test with a general to specific procedure starting with 20 lags using the t-statistic on the last lag. We have estimated all models both with trend and constant, with constant and without trend or constant. However, neither the joint test of a unit root and no linear trend nor the joint test of a unit root and no constant was acceptable. For the test in levels, the test statistic for the former equalled 3.93 (10% critical value is 5.47) while the test statistic for the hypothesis of a unit root but no constant equalled 1.86 (10% critical value: 3.86). The test for the first differences was 4.27 (unit root but no trend) and 4.24 (unit root but no constant). Hence Table 2 only reports ADF-tests for the model without trend and constant. Although the ADF-test reveals an explosive process for p_t, the unit root hypothesis is rejected for Δp_t.

Our second test is the PP-test. The table only reports the results for the test without a trend but the conclusions are not affected if a trend is added. The PP-test clearly reveals that the series p_t is difference stationary. Third, we have done a KPSS-test to confirm our findings from the ADF and PP-tests. The null of stationary series is clearly rejected for p_t but the test is unable to reject this hypothesis for Δp_t.

Because the series possibly exhibits a break in 1998, we have used the Perron (1997) test to check if the series contain a unit root if this structural break is accounted for. We used the procedure that minimizes the t-statistic on θ to select the break date. With the exception of the estimates of Eq. (7) for Δp_t, all endogenously determined breaks are located between February 1998 and May 1998. The estimates for p_t further reveal positive significant values of ψ in Eq. (5) and φ in Eq. (7). The estimates of δ in Eqs. (5) and (6) are also positive and significant for Δp_t. The significance of these values notwithstanding, Table 2 supports the hypothesis that the series p_t contains a unit root and the series Δp_t is stationary.

Table 2
Unit root tests for p_t and Δp_t

Variable	Test	Test statistic	Lags[a]
p_t	ADF	0.2006	9
p_t	PP (constant)	−1.4337	4
p_t	KPSS[b] (no trend)	0.1739***	4
p_t	KPSS[b] (trend)	0.8389**	4
p_t	PP − Eq. (5)	−4.4863	7
p_t	PP − Eq. (6)	−4.6832	11
p_t	PP − Eq. (7)	−3.2258	7
Δp_t	ADF	−2.9194***	8
Δp_t	PP	−7.9048***	4
Δp_t	KPSS[b] (no trend)	0.1315	4
Δp_t	KPSS[b] (trend)	0.0935	4
Δp_t	PP − Eq. (5)	−9.0228***	0
Δp_t	PP − Eq. (6)	−9.0366***	0
Δp_t	PP − Eq. (7)	−3.3899	8

*, ** and *** refer to significance at the 10%, 5% and 1% levels.

[a] Lags refer to the number of lags obtained following the general to specific procedure for the ADF and endogenous break (PP models) and to the number of lags used for the KPSS and PP-tests.

[b] KPSS 1% critical value without (with) trend equals 0.739 (0.216), the 5% critical value without (with) trend equals 0.463 (0.146).

Based on the evidence from the ADF, PP and PP with endogenously determined time breaks, the hypothesis that the SO_2 permit price series contains a unit root cannot be rejected. For the first differenced series, on the other hand, the evidence clearly suggests that the hypothesis of a unit root should be rejected. The KPSS-test does not allow us to accept the null of stationarity for p_t while it fails to reject the null of stationarity for Δp_t. All in all, this suggests that p_t is a non-stationary process. However, this is not sufficient to conclude that the series is a random walk.

3.3. Tests of the random walk hypothesis

Although the unit root tests have clearly shown that permit prices contain a unit root, this is not sufficient to adopt the random walk hypothesis (Campbell et al., 1997). Indeed, the various random walk hypotheses impose restrictions on the error process that have not been analysed so far. We will start with the restriction imposed by the RW3 model. For the RW3 model, condition (2) requires that all autocorrelations between Δp_t and Δp_{t-k}, $\rho(k)$ equal zero for all values of $k > 0$. Table 3 provides estimates of the autocorrelation coefficients for the level (p_t), Δp_t and $(\Delta p_t)^2$. The Ljung–Box $Q(k)$ statistic allows to assess the significance of these coefficients. The results for p_t reveal a typical pattern for a non-stationary series: the autocorrelation coefficient for $k = 1$ is close to unity and dies out slowly (Enders, 1995). With respect to Δp_t, the results suggest that $E[\Delta p_t \, \Delta p_{t-k}] \neq 0$ for various lag lengths. However, the correlation coefficients seem to be small and are only significant at the 5% level until k reaches 5. The significance of the autocorrelation coefficients is evidence against the hypothesis that the series p_t is a random walk of type 3. The evidence with respect to the autocorrelation coefficients for $(\Delta p_t)^2$ suggests that the series for Δp_t exhibits volatility clustering as $E[(\Delta p_t)^2 \, (\Delta p_{t-k})^2] \neq 0$ implies that $E[\sigma_t^2, \sigma_{t-k}^2] \neq 0$. The estimated coefficients are, however, small and are no longer significant at a 5% level for values of $k > 5$. Furthermore, one should be very careful with respect to the $Q(k)$ statistic as it is a joint test on all autocorrelations up to a certain level k. Hence, significance (say at $k = 5$) could be due to one strongly significant autocorrelation (for instance, at $k = 2$).

A second test on Δp_t is based on the ratio of the variances at two different frequencies. Lets assume that we

Table 3
Autocorrelations for p_t, Δp_t and $(\Delta p_t)^2$

Lag k	p_t		Δp_t		$(\Delta p_t)^2$	
	Autocor.	$Q(k)$	Autocor.	$Q(k)$	Autocor.	$Q(k)$
1	0.967	86.20	0.180	2.96*	−0.011	0.01
2	0.924	165.67	0.268	9.60***	0.344	10.91***
3	0.862	235.76	−0.046	9.81**	0.075	11.44***
4	0.803	297.33	−0.003	9.81**	−0.056	11.75**
5	0.744	350.82	−0.050	10.05*	0.012	11.76**
6	0.688	397.07	−0.023	10.10	−0.034	11.88*
7	0.633	436.68	0.181	13.31*	−0.049	12.12*
8	0.566	468.71	0.054	13.61*	−0.045	12.32
9	0.494	493.49	−0.081	14.27	−0.079	12.95
10	0.428	512.32	−0.047	14.50	−0.073	13.50
11	0.365	526.16	−0.072	15.03	−0.034	13.62
12	0.306	536.03	−0.054	15.35	−0.071	14.16
13	0.251	542.76	−0.121	16.91	0.095	15.13
14	0.204	547.27	−0.064	17.35	−0.076	15.76
15	0.161	550.12	0.014	17.37	0.041	15.95
16	0.117	551.66	0.018	17.41	0.004	15.95
17	0.073	552.26	−0.027	17.49	−0.059	16.34
18	0.029	552.37	−0.133	19.52	−0.059	16.74

*, ** and *** refer to significance at the 10%, 5% and 1% levels.

compare Δp_t values at a monthly interval and a larger interval of q months. If we were to set $q = 3$, for instance, this would mean that we would compare monthly and quarterly changes in the permit price. The former equals Δp_t while the return at an interval equal to q equals $\sum_{i=0}^{q-1} \Delta p_{t-i}$. Because of Eq. (2), the variance of the latter equals q times the variance of the former. Campbell et al. (1997) derive a test statistic under RW3 that allows testing if the variance of the return at the q-interval is equal to q times the return at the monthly interval. The test statistic is derived from the sample autocorrelations and is given by:

$$\psi(q) = \frac{\sqrt{T-1}(\mathrm{VR}(q)-1)}{\sqrt{\hat{\theta}}} \sim N(0,1) \tag{8}$$

with

$$\mathrm{VR}(q) = 1 + 2\sum_{k=1}^{q-1}\left(1-\frac{k}{q}\right)\hat{\rho}(k)$$

and $\hat{\theta}(q)$, a heteroskedasticity-consistent estimator of the variance of $\mathrm{VR}(q)$ (for details, see Campbell et al., 1997, p. 55).

The evidence presented in Table 4 seems to reinforce the conclusions in Table 3. In line with the evidence presented in Table 3, the fact that $\mathrm{VR}(q) > 1$ implies that the autocorrelations are positive. Furthermore, the variance ratio test rejects the hypothesis that the permit price series is a random walk of type 3 for levels of $q < 6$ at a 5% level of significance. However, for levels of $q \geq 6$, which compare monthly returns to, for instance, yearly returns, the test fails to reject the null of no significant autocorrelation among the returns at the 5% level.

Both the evidence from the autocorrelation coefficients and the variance ratio tests offer some support for the hypothesis that the permit price process is not a random walk of type 3. The evidence presented here is not all that different from the evidence for financial markets. Lo and MacKinlay (1999) for instance find that autocorrelations are not all zero. The evidence against the hypothesis that the series is a random walk of type 3 is, however, not overwhelming. The autocorrelation coefficients presented in Table 3 for instance are small and the variance ratio fails to reject the RW3 hypothesis for levels of $q \geq 6$. The size of the autocorrelations coefficients would suggest that one can question whether the significance in a statistical sense extends to significance in an economic sense.

3.4. Predictability

The question that emerges from the previous paragraph is whether the significant autocorrelations can

Table 4
Variance ratio test

q	$\mathrm{VR}(q)$	$\psi(q)$	Sig. $\psi(q)$
3	1.4188	2.2401**	0.0251
4	1.5137	2.0299**	0.0424
5	1.5703	1.8931*	0.0583
6	1.5913	1.7470*	0.0806
7	1.5998	1.6251	0.1041
8	1.6501	1.6457*	0.0998
9	1.7004	1.6773*	0.0935
10	1.7261	1.6600*	0.0969
11	1.7384	1.6221	0.1048
12	1.7375	1.5645	0.1177

*, ** and *** refer to significance at the 10%, 5% and 1% levels.

be exploited from an economic point of view. Market participants have to incur transaction costs if they buy or sell a permit. If autocorrelations can be exploited to earn a profit, one cannot argue that SO_2 permit markets are efficient. New information which arrives and has an impact on the value of an SO_2 permit is not immediately reflected in its price. Hence, from an informational efficiency point of view, this would be evidence against efficient markets. If one cannot earn such a profit it follows that all information that affects the value of SO_2 permits is included in the permit price. However, given the statistical significance of autocorrelations, new information is only reflected in prices up to such a level where it is possible to profit from the price history. Hence, from an economic point of view, exploiting the significant autocorrelations fully is not rational if one takes into account transaction costs and one can argue that permit prices reflect all information which is significant from an economic point of view.

It follows that the issue that needs to be addressed is whether the SO_2 permit price history can be used to earn a profit. Obviously, there are various ways to test this hypothesis. We have chosen to estimate a time series model which allows us to predict next period's change and to see whether it could have been used to predict permit prices with a relative high level of certainty. In order to do so, we have estimated an ARMA model for the first difference of the permit price series. As was shown in the previous section and in Table 3, $\mathrm{E}[(\Delta p_t)^2\,(\Delta p_{t-k})^2] = \mathrm{E}\left[\sigma_t^2, \sigma_{t-k}^2\right] \neq 0$. This suggests that the price process exhibits time-varying volatility. In order to capture this property of this process, one can use a member of the large family of autoregressive conditional heteroskedasticity (ARCH) models (see for instance Chan et al., 2005a,b). After some experimentation with various ARMA–GARCH models we found that the model which best serves out purposes was the AR(2)–GARCH(1,1) model:

$$\Delta p_t = \delta_1 \Delta p_{t-1} + \delta_2 \Delta p_{t-2} + \xi_t \tag{9}$$

with $\xi_t = \eta_t \sqrt{\sigma_t}$, $\eta_t \sim \mathrm{IID}(0,1)$ and

$$\sigma_t^2 = \omega + \alpha \xi_{t-1}^2 + \beta \sigma_{t-1}^2 \tag{10}$$

with ω as estimate of the long run variance.

The model requires that $|\delta_i| < 1$, $i = 1,2$; $\omega > 0$; $\alpha \geq 0$ and $\beta_i \geq 0$ (Bollerslev, 1986). Table 5 reports the results. They suggest that the AR(2)–GARCH(1,1) model is appropriate. From the Q-statistics for the residuals, it can be seen that they do not exhibit any autocorrelation which suggests that the estimates capture all of the autocorrelation. Secondly, the $Q(k)$ statistics for the squared residuals reveal that the model removes most of the autocorrelation as only the $Q(3)$ statistic is significant at the 10% level of significance. Furthermore, the ARCH-LM test does not allow dismissing the null hypothesis of no ARCH-effects. This suggests that the GARCH part of the model captures the time-varying volatility. The estimate of ω in Table 5 (0.000714) is not statistically different from the variance presented in Table 1 (0.0004). This suggests that the AR(2)–GARCH(1,1), as far as forecasting is concerned, is an appropriate model as the residuals nor their square contain any information which could be used for forecasting purposes.

Table 5
AR(2)–GARCH(1,1) estimates

	Estimate	Standard error[a]
AR(2)		
Δp_{t-1}	0.212831*	0.122596
Δp_{t-2}	0.218871	0.168118
GARCH(1,1)		
ω	0.000714**	0.000356
ξ_{t-1}^2	0.357813*	0.193965
σ_{t-1}^2	0.587051***	0.154951
Uncentred R^2	0.0846	
Residuals		
$Q(3)$	0.6913	
$Q(4)$	0.8000	
$Q(5)$	1.7584	
Squared Res.		
$Q(3)$	3.6226*	
$Q(4)$	4.2237	
$Q(5)$	4.3018	

*, ** and *** refer to significance at the 10%, 5% and 1% levels.
[a] Standard errors are heteroskedasticity-consistent (Bollerslev–Wooldridge).

The results in Table 5 suggest that it would be hard to make a profit based on the past information. Although the estimates are significant, the R^2 is very low which is indicative of the fact that the model is not able to predict future SO_2 permit prices with much certainty. As a matter of fact only 8% of the variation in the permit price process is explained. Hence, one can question if it would be possible to profit from knowledge of price history on the SO_2 permit market. As was the case with the non-zero autocorrelations coefficients, this feature is also present in financial markets (Malkiel, 2003). The estimates of the GARCH-terms confirm that the variance in SO_2 permit markets clusters.

This suggests that it would have been impossible to exploit the significant autocorrelations from an economic point of view. Market participants react fast to new information and continue to do so up until the point where it is no longer possible for them to use the information in their market behaviour. This clearly suggests that the US SO_2 permit market is efficient and that SO_2 permit prices reflect all information which are significant from an economic point of view. Market participants have a good understanding of the price process. If this was not the case, returns would be predictable and one would be able to exploit the price process to earn a profit. Assume for instance that market participants tend to overreact and that news arrives which causes the market price to jump upwards. To the extent that market participants overreact today, we would expect a correction (negative return) in the future. Hence, permit prices would be predictable: a spike would be followed by a correction. If these overreactions are large enough, our model would be able to explain a large part of the variation in the price produces. Furthermore, knowledge of the model would allow one to earn a profit after transaction costs by buying (selling) a permit today and selling (buying) the same permit in the following period. Knowledge of the model would indeed allow a participant to predict with reasonable certainty next period's prices.

If, on the other hand, market participants are slow to adjust prices, permit prices would exhibit a series of positive or negative returns. Hence they would be predictable. Again, this would imply that our model is able to explain a large part of the variation.

Our model, although it seems reasonable from a statistical point of view, could hardly be used to predict with a reasonable amount of certainty next period's permit prices. The lack of a reasonable amount of certainty in terms of prediction couples with the existence of transaction costs clearly suggests that past price behaviour is not informative. Hence, from an economic point of view, permit markets are efficient.

4. Discussion and conclusion

The evidence presented in this paper suggests that the market for SO_2 permits in the US is not all that different from financial markets. For financial markets, the random walk hypothesis (RW3) is also often rejected. However, as is the case for the SO_2 permits market; economic profitable predictability is mostly rejected as well. Hence, although one cannot reject the hypothesis that this market is weakly efficient from a statistical point of view, the economic significance of the predictability is very limited if not nonexistent.

The evidence presented in this paper suggests that new information is reflected in the permit price fast. The SO_2 permit market is basically as efficient as financial markets. This is clearly important as it is indicative of the fact that the value of SO_2 permits reflects all relevant information. This suggests that market participants have a good understanding of the price process and have a good understanding of the way in which new information affects the market-clearing price. The evidence presented in this paper supports the conclusion in Joskow et al. (1998) that it would be

> "*hard to argue that bidders in the 1993 auctions had a good idea of a single market-clearing price. It would be a good deal easier to make this argument for the 1994 auctions*." (p. 681).

Based on our analysis of the history of SO_2 permit prices, we reach the same conclusion from a different perspective.

New information that increases the variance has the tendency to cluster. If an event increases uncertainty, this uncertainty does not return back to its previous level in one period. The significant GARCH-effects suggest that it takes time before it settles down.

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